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Implementing Second-Best Environmental Policy under Adverse Selection

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Implementing second-best environmental policy under adverse selection

Abstract

A key obstacle to practical application of mechanism design theory in regulation is the difficulty of obtaining consistent beliefs regarding information assumed to be commonly held in the models. This paper presents a solution to this problem by developing an easily-implemented empirical methodology with which the government can use commonly available data to develop beliefs regarding the technology and distribution of types in a regulated sector characterized by hidden information. Results are used to calibrate a second-best land conservation mechanism and evaluate its cost relative to simpler alternatives.

Key words: Empirical contract theory, environment, conservation reserve program, stochastic frontier analysis

1 Introduction

The normative theoretical literature on optimal regulation under adverse selection has grown tremendously in the past three decades. In spite of this progress, actual policies implementing even the most basic optimal mechanisms remain scarce. One obstacle to the transition from theory to practice is the difficulty of obtaining information that the theoretical models assume to be commonly held. Models typically characterize a second-best (as opposed to the full information first-best) menu of contracts stipulating payments and allocations among which the regulated firms choose. The precise terms of each contract are defined up to commonly held beliefs regarding: a) firms' production technology; and b) the probability distribution of firm types.[1] The optimal values of the contract terms can vary greatly depending on these two sets of beliefs. From the standpoint of applied theory, the development of consistent beliefs regarding these items is therefore of paramount importance. How to develop these beliefs given readily available information (e.g., from industrial surveys) is an issue on which the literature has largely remained silent.

This paper shows how to calculate the terms of a second-best mechanism by modeling type as a source of heterogeneity that is unobserved both to the regulator and the econometrician. Such a methodology has a wide range of potential applications for general principal-agent problems characterized by adverse selection. It is particularly relevant in regulatory applications that fit into the general framework of Baron and Myerson (1982). With respect to environmental policy, this framework readily applies to cases in which environmental benefits are privately provided or in which firms have an effective right (e.g., for legal or political reasons) to pollute.

As a concrete application, I consider the problem of designing a program to encourage

[1] Type refers to a productivity parameter that is private information to each firm.

landowners to set aside environmentally-sensitive land from production. Such conservation payments programs are becoming popular in both the developed and developing world. Prominent examples include the Wetlands Reserve Program and Conservation Reserve Program (CRP) in the United States and Costa Rica's Pago de Servicios Ambientales (payment for environmental services) program. Such programs share the characteristics that participation is voluntary, and that opportunity costs of participation (i.e., foregone profit) are both heterogeneous across landowners and not directly observable to the government.

There are several policy instruments available to the government for meeting a given land set aside target. The choice of instrument involves a trade-off between simplicity and cost-effectiveness. At one extreme is a single Pigouvian subsidy. This instrument is simple to administer, but involves potentially large excess payments to landowners with low participation costs. At the other extreme is first-degree price discrimination. This instrument is difficult to implement. It requires the government to obtain perfect information regarding costs and to design a contract tailored to each landowner. It is cost effective, however, since it results in an efficient allocation with each owner receiving a payment exactly equaling his opportunity cost. Between these extremes lie third-degree price discrimination (geographically differentiated Pigouvian subsidies) and second-degree price discrimination (the Baron and Myerson mechanism).

Theory can rank these instruments by expected cost. With the empirical tools provided in this paper, however, one can go a step further towards making an informed choice between instruments based on the magnitude of cost differences. Such analysis also contributes to the discussion in the mechanism design literature surrounding the "Wilson doctrine." In an influential piece, Wilson (1987) noted the prevalence of simple rather than theoretically optimal trading rules in markets with asymmetric information. As noted above, the terms of a second-best contract mechanism are highly dependent upon the regulator's beliefs regarding the production technology and distribution of types. In contrast, some mechanisms (like Pigouvian instruments and some types of auctions) are always allocatively efficient, even if they do not optimize the regulator's objective. Proponents of the Wilson doctrine argue that such mechanisms (which also have the virtue of simplicity) are therefore preferable to a complicated mechanism that is optimal only if the regulator's beliefs are correct. The analysis illustrated here provides policy-makers an easily implementable means of evaluating the magnitude of potential gains offered by theoretically optimal mechanisms.

To identify both the production technology and the distribution of types, the empirical methodology uses a two-part additive error structure in the spirit of the stochastic frontier models pioneered by Aigner et al. (1977) and Meeusen and van den Broeck (1977). One part of the error is stochastic noise, while the other represents type. This econometric model extends the stochastic frontier literature by adapting robust generalized method of moments (GMM) estimation techniques recently developed in other contexts.

Previous research attempting to identify the production technology and distribution of types under adverse selection can be divided into two branches. One branch of the empirical contract theory literature, such as Wolak (1994), Thomas (1995), and Lavergne and Thomas (2005) assumes that existing regulations are optimal in a second-best sense. The optimality conditions provide equations that can be econometrically estimated with observable data. Regression results allow the econometrician to infer the regulator's beliefs regarding the distribution of types and the technology. By its nature, this line of research is descriptive rather than prescriptive. It may describe the regulator's beliefs regarding the commonly held information. However, it provides no guidance to the regulator as to how she might develop them in the first place.

A second branch of the literature deals with how to estimate the actual distribution of types rather than the regulator's beliefs. In initial work along this line, Dalen and Gómez-Lobo (1997) interpret regression residuals as firm types for a model of urban transport. The model is rather restrictive since it does not allow for any random error. A more flexible approach allows for both random types and stochastic noise. Bousquet and Ivaldi (1997) and Gagnepain and Ivaldi (2002) use such an approach in their studies of French telecommunications and urban transport. They assume parametric distributions for both unobserved variables. Maximum likelihood techniques allow estimation of the parameters of the relevant value functions and probability distributions.

By allowing for both random error and unobserved types, Bousquet and Ivaldi (1997) and Gagnepain and Ivaldi (2002) are the works most closely related to this paper. The methodology presented here has three principal advantages over earlier approaches. First, as pointed out by Kopp and Mullahy (1990), a disadvantage to using the maximum likelihood approach often employed in stochastic frontier analysis is that parameter estimates are inconsistent in the presence of errors that are not i.i.d. The approach developed here is robust to arbitrary heteroskedasticity and geographic clustering. Second, these papers estimated a single equation. In a framework with two unobserved sources of variation, deriving and estimating the likelihood function for a joint system of several equations with potentially correlated errors becomes quite cumbersome, and to my knowledge has not been done.[2] In contrast, the framework employed here easily allows one to increase estimation efficiency by using a system of cost function and expenditure share equations. Finally, this approach is easily implementable with cross-sectional data, and is computationally undemanding.

[2] Current techniques for maximum likelihood estimation of stochastic frontier systems require the assumption that the error term in the cost frontier equation is independent of the error terms in the expenditure share equations for the same observation (see Kumbhakar and Lovell, 2000).

The rest of the paper is organized as follows. In the next section, I create a formal theoretical model of the four policy options for setting aside land. In Section 3, I develop and implement the econometric strategy for obtaining consistent beliefs regarding the production technology and distribution of agent types for a sample of Midwestern U.S. farmers. In Section 4, I use this information to calibrate a simulation evaluating the relative costs of the various land set aside policies. Section 5 contains concluding comments.

2 Alternative Contract Schemes

I consider four versions of a program designed to induce landowners to remove land from production. In all versions, the program has two salient features. First, the government must ensure that the sector idles a targeted quantity of land. Second, the program must be voluntary. The regulator's objective is to allocate set asides and transfer payments to each landowner that minimize the expected cost of satisfying the constraints.

The basic theoretical model builds upon earlier work by Smith (1995) and Crépin (2005). Landowners can be differentiated by observable characteristics that depend on a general geographical location (referred to as their *county*), indexed by $i \in \{1, ..., I\}$. In a given county landowners are assumed to be identical up to an unobserved (to the government) productivity parameter $\theta \in (0, 1]$, referred to as their type. Let $f(\theta) \equiv dF(\theta)/d\theta$ denote the government's beliefs regarding the probability density function of types, assumed to be the same for all counties. This function satisfies the monotone hazard rate condition: $d[F(\theta)/f(\theta)]/d\theta > 0$.[3]

Let a denote the amount of land idled by an individual. The restricted profit function

[3] This assumption is standard in mechanism design theory. It is satisfied by a wide range of common distributions (see Bagnoli and Bergstrom, 2005).

$r^i(a,\theta)$, indicates the market income obtained from the unenrolled land.[4] It completely characterizes the production technology. This function satisfies $r^i_\theta > 0$, $r^i_a < 0$, $r^i_{aa} < 0, r^i_{a\theta} < 0$, and $r^i_{aa\theta} < 0$.[5] The first property indicates that θ increases productivity. The next indicates that enrolling land reduces profit. The third property shows that this marginal opportunity cost becomes greater in absolute value as enrollment increases (least productive land is enrolled first). The fourth property, commonly referred to as the single-crossing condition, indicates that the marginal opportunity cost is increasing in type (all else equal, more productive landowners forego more profit from idling an additional acre). The final property is a regularity condition that helps ensure that the second-best problem is well behaved.

The pair $\langle a^i(\theta), t^i(\theta)\rangle$ denotes contract terms in county i for type θ, where $a^i(\theta)$ is the amount of land enrolled in the program and $t^i(\theta)$ is the transfer payment. The environmental constraint is modeled as a requirement that the contracts meet the enrollment target:

$$\sum_{i=1}^{I} n^i \int_0^1 a^i(\theta)\, dF(\theta) \geq A \sum_{i=1}^{I} n^i. \tag{1}$$

Here, A is a parameter indicating the average acreage idled per landowner, and n^i is the number of landowners in county i. Maximum enrollment for any individual is constrained by his total available acreage

$$a^i(\theta) \leq \bar{a}^i \;\forall\, \theta, i, \tag{2}$$

where $\bar{a}^i$ is the size of a holding in county i. Let $\lambda \geq 0$ and $\gamma(\theta) \geq 0$ respectively denote the Lagrange multipliers for (1) and (2).

[4] Output price is assumed to be perfectly elastic so that a landowner's profit is not affected by the amount of land idled by others.

[5] As a notational convention, subscripts on functions denote partial derivatives and superscripts denote county indexes.

The second policy constraint is that program participation be voluntary. As a result, the regulator must compensate landowners for at least their opportunity cost:

$$t^i(\theta) \geq r^i(0, \theta) - r^i\left(a^i(\theta), \theta\right), \quad \forall\, \theta, i. \tag{3}$$

Define surplus payments received in excess of the minimum necessary to satisfy (3) by:

$$s^i(\theta) \equiv t^i(\theta) - \left[r^i(0, \theta) - r^i\left(a^i(\theta), \theta\right)\right]. \tag{4}$$

The voluntary participation constraint (3) can then be expressed more succinctly as:

$$s^i(\theta) \geq 0, \quad \forall\, \theta, i. \tag{5}$$

It is convenient to use Eq. (4) to express expected program cost in terms of surplus rather than transfers:

$$\sum_{i=1}^{I} n^i \int_0^1 \left[s^i(\theta) + r^i(0, \theta) - r^i\left(a^i(\theta), \theta\right)\right] dF(\theta). \tag{6}$$

The policy alternatives in the following subsections progress from least to most constrained. The first-best allows the government to allocate transfers and payments across types as it sees fit, subject only to constraints (1), (2), and (5). The second-best adds the constraint that the allocation must be incentive compatible. The county-level Pigouvian subsidy requires that the same per-acre payment be offered to all landowners in a given county, with participation levels determined by each individual. Since this system is necessarily incentive compatible, it is more constrained than the second-best. The last alternative adds the constraint that a single per-acre payment be offered to all landowners, regardless of location. Therefore, the cost of attaining the land set aside target cannot be decreasing as we progress from one alternative to the next. The magnitude of any possible increase in cost is the empirical question to be addressed in Sections 3 and 4.

2.1 First-Best (First-Degree Price Discrimination)

The Lagrangian for the first-best minimization problem can be expressed:

$$\begin{aligned} L^{FB} &= \sum_{i=1}^{I} n^i \int_0^1 \left\{ s^i(\theta) + r^i(0,\theta) - r^i\left(a^i(\theta),\theta\right) \right. \\ &\quad \left. + \lambda\left[A - a^i(\theta)\right] - \tfrac{\gamma^i(\theta)}{f(\theta)}\left[\bar{a}^i - a^i(\theta)\right] \right\} dF(\theta) \end{aligned} \tag{7}$$

subject to surplus constraint (5). Since surplus payments only increase cost, this constraint binds for all types. The first-best land allocation also satisfies the following Kuhn-Tucker conditions for all i and θ:

$$r_a^i\left(a^i(\theta),\theta\right) + \lambda - \frac{\gamma^i(\theta)}{f(\theta)} \leq 0; \tag{8}$$

$$a^i(\theta)\left[r_a^i\left(a^i(\theta),\theta\right) + \lambda - \frac{\gamma^i(\theta)}{f(\theta)}\right] = 0; \tag{9}$$

$$\lambda \int_{\Theta} \left[A - a^i(\theta)\right] dF(\theta) = \gamma^i(\theta)\left[\bar{a}^i - a^i(\theta)\right] = 0. \tag{10}$$

Consequently, for an interior solution in which neither county nor individual acreage restrictions bind, the optimal first-best program satisfies the equimarginal principle. It equates the marginal profit from cultivating an additional acre of land for each landowner to the shadow cost of tightening the environmental constraint for the entire sector. The first-best program could be implemented with a two-part tariff. The regulator would offer a uniform subsidy for idled land equal to the shadow value of the enrollment target. Landowners would respond by idling the efficient quantity of land. A type-dependent lump-sum tax would recover all surplus payments arising from the subsidy.

2.2 Second-Best (Second-Degree Price Discrimination)

The optimal second-best policy is slightly more complicated. The Revelation Principle (e.g., Myerson, 1979) indicates that there is no loss in generality by restricting attention to direct revelation mechanisms satisfying incentive compatibility. Incentive compatibility requires that for all i:

$$\theta \in \arg\max_{\tilde{\theta}} \left\{ r^i \left(a^i \left(\tilde{\theta} \right), \theta \right) + t^i \left(\tilde{\theta} \right) \right\} \ \forall \ \left(\theta, \tilde{\theta} \right) \in \Theta^2. \tag{11}$$

This requirement, combined with the participation constraint (5), imposes two restrictions on the set of feasible contract allocations (both follow directly from results in Baron and Myerson (1982)). First, for an interior solution, enrollment is monotonically non-increasing in type:

$$\frac{da^i(\theta)}{d\theta} \leq 0. \tag{12}$$

Second, expected surplus is weakly decreasing over type at the rate

$$\frac{ds^i(\theta)}{d\theta} = r^i_\theta \left(a^i(\theta), \theta \right) - r^i_\theta (0, \theta). \tag{13}$$

Intuitively, if the government offered payments to exactly offset opportunity costs, low types would choose contracts designed for higher types (but not vice versa). Therefore, the lower the type, the higher the surplus payment necessary to induce a landowner to choose the contract intended for him.

Since surplus is non-increasing, the best the principal can do while satisfying (5) and (13) is to set $s^i\left(\bar{\theta}\right)=0$. Using Eq. (13), surplus is then

$$s^i\left(\theta\right)=\int_{\theta}^{1}\left[r_{\theta}^i\left(0,\omega\right)-r_{\theta}^i\left(a^i\left(\omega\right),\omega\right)\right]d\omega. \tag{14}$$

Temporarily ignoring (12), substitution of Eq. (14) into Eq. (6) and integrating by parts yields the following Lagrangian for the government's second-best problem:

$$\begin{aligned} L^{SB} = & \sum_{i=1}^{I} n^i \int_0^1 \left\{ \tfrac{F(\theta)}{f(\theta)} \left[r_{\theta}^i\left(0,\theta\right)-r_{\theta}^i\left(a^i\left(\theta\right),\theta\right)\right]+r^i\left(0,\theta\right)-r^i\left(a^i\left(\theta\right),\theta\right) \right. \\ & \left. +\lambda\left[A-a^i\left(\theta\right)\right]-\tfrac{\gamma^i(\theta)}{f(\theta)}\left[\bar{a}^i-a^i\left(\theta\right)\right]\right\} dF\left(\theta\right). \end{aligned} \tag{15}$$

The second-best land allocation satisfies the following Kuhn-Tucker conditions:

$$r_a^i\left(a^i\left(\theta\right),\theta\right)+\tfrac{F(\theta)}{f(\theta)}r_{a\theta}^i\left(a^i\left(\theta\right),\theta\right)+\lambda-\tfrac{\gamma^i(\theta)}{f(\theta)}\leq 0; \tag{16}$$

$$a^i\left(\theta\right)\left[r_a^i\left(a^i\left(\theta\right),\theta\right)+\tfrac{F(\theta)}{f(\theta)}r_{a\theta}^i\left(a^i\left(\theta\right),\theta\right)+\lambda-\tfrac{\gamma^i(\theta)}{f(\theta)}\right]=0; \tag{17}$$

$$\lambda\int_{\Theta}\left[A-a^i\left(\theta\right)\right]dF\left(\theta\right)=\gamma^i\left(\theta\right)\left[\bar{a}^i-a^i\left(\theta\right)\right]=0. \tag{18}$$

The restrictions on $F\left(\theta\right)$ and $r^i\left(\cdot\right)$ ensure that this solution satisfies (12) (see Guesnerie and Laffont, 1984).

The impact of asymmetric information can be easily seen for interior solutions. Unlike the first-best case, rather than having the marginal profit of land be equated for all farms, there is a distortion created by the term $F\left(\theta\right)r_{a\theta}^i/f\left(\theta\right)$ in Eq. (16). As a result, the equimarginal principal is never satisfied. This program could be implemented by the government requesting that landowners choose a contract from a menu of possible choices. Alternatively, the second-best welfare level can be thought of as an upper bound of the welfare obtainable by an optimally designed procurement auction. For an interior solution, the second-best is not obtainable by a Pigouvian subsidy.

2.3 Differentiated Pigouvian Subsidy (Third-Degree Price Discrimination)

This program is also geographically differentiated inasmuch as a distinct payment scheme can be applied to each county. Unlike the previous case, however, transfers are restricted to be the product of per-acre Pigouvian subsidy τ^i and enrolled acres (i.e., $t^i(\theta) \equiv \tau^i a^i(\theta)$). The first order condition of incentive compatibility condition (11) requires that an interior solution satisfy:

$$-r_a^i\left(a^i(\theta), \theta\right) = \tau^i. \tag{19}$$

Presented with this subsidy, the solution the landowners' optimal enrollment problem is:

$$\tilde{a}^i(\tau, \theta) \equiv \arg\max_a \left\{r^i(a, \theta) + \tau^i a\right\}.$$

Knowing this, the government's optimization problem is to choose the vector $\boldsymbol{\tau} \equiv \left(\tau^1, ..., \tau^I\right)$ $\in \Re_+^I$ that minimizes expected expenditures. The corresponding Lagrangian is:

$$L^{CP} = \sum_{i=1}^{I} n^i \int_0^1 \left\{\tau^i \tilde{a}^i\left(\tau^i, \theta\right) + \lambda\left[A - \tilde{a}^i\left(\tau^i, \theta\right)\right] - \tfrac{\gamma^i(\theta)}{f(\theta)}\left[\bar{a}^i - \tilde{a}^i\left(\tau^i, \theta\right)\right] dF(\theta)\right\}. \tag{20}$$

The optimal solution satisfies the following conditions:[6]

$$\int_0^1 \tilde{a}^i\left(\tau^i, \theta\right) dF(\theta) + \tau^i \tfrac{\partial \int_0^1 \tilde{a}^i\left(\tau^i, \theta\right) dF(\theta)}{\partial \tau^i} \geq \tag{21}$$
$$\lambda \tfrac{\partial \int_0^1 \tilde{a}^i\left(\tau^i, \theta\right) dF(\theta)}{\partial \tau^i} - \tfrac{\partial \int_0^1 \frac{\gamma^i(\theta)}{f(\theta)} \tilde{a}^i\left(\tau^i, \theta\right) dF(\theta)}{\partial \tau^i};$$

[6]As in the standard monopsonist problem, it is necessary to verify satisfaction of the second order condition since the restrictions imposed on $r(\cdot)$ are not sufficient to guarantee a minimum.

$$\lambda \int_0^1 \left[A - \tilde{a}^i \left(\tau^i, \theta\right)\right] dF\left(\theta\right) = \gamma^i\left(\theta\right) \left[\bar{a}^i - \tilde{a}^i \left(\tau^i, \theta\right)\right] = 0. \tag{22}$$

The intuition behind this result is easiest understood for interior solutions in which the optimal τ^i is strictly positive and constraint (2) does not bind. Then, Eq. (21) can be rearranged to yield a variant of the inverse-elasticity pricing rule for a monopsonist that can discriminate between distinct markets:

$$1 + \frac{1}{\varepsilon^i} = \frac{\lambda}{\tau^i}. \tag{23}$$

Here,

$$\varepsilon^i \equiv \frac{\partial \int_0^1 \tilde{a}^i\left(\tau^i, \theta\right) dF\left(\theta\right)}{\partial \tau^i} \cdot \frac{\tau^i}{\int_0^1 \tilde{a}^i\left(\tau^i, \theta\right) dF\left(\theta\right)} \tag{24}$$

is the price elasticity of expected supply of enrolled acreage in county i. The government minimizes cost by providing a higher subsidy in counties with more elastic supply.

2.4 Single Pigouvian Subsidy (No Price Discrimination)

The single Pigouvian subsidy is similar to the case analyzed in the previous section, with the exception that all τ^i must share the same value. Therefore, when landowners respond to the subsidy, for an interior solution the Pigouvian land allocation equates marginal returns to land (just like the first best). Moreover, in order to satisfy the environmental constraint (1), this land allocation must be exactly the same as for the first-best program (i.e., equal to the shadow value of enrolled land). Unlike the first-best, however, the information asymmetry prevents the use of a two-part tariff to recover surplus payments. The best it can do is ensure that the highest participating type receive zero surplus.

3 Empirical Model

In this section, I show how one can estimate the parameters of the production technology and the distribution function of agent types using commonly available data (e.g., from cross-sectional industrial surveys). These results provide the necessary information to calibrate the allocations for the policy alternatives described above. I begin by specifying a parametric technology. This technology implies a cost function, expenditure share equations (by Shephard's Lemma), and a revenue to cost ratio (by profit maximization). I then show how robust GMM techniques can be applied to standard stochastic frontier analysis for efficient estimation of this system of equations.

3.1 Specification

Since the Aigner et al. (1977) and Meeusen and van den Broeck (1977) stochastic frontier models employ log-linear econometric models, the presence of negative observed profit in the data makes it is more convenient to work with the cost function rather than the restricted profit function directly. Once the parameters of the cost function are estimated, one can use them to recover the profit function. Let $C(\mathbf{w}, q, \ell, \theta)$ denote the minimum cost for a type θ to produce $q \in \Re_+$ units of an aggregate agricultural commodity, given a vector of variable input prices $\mathbf{w} \equiv (w_1, ..., w_N) \in \Re^N_{++}$ and a land endowment $\ell \in \Re_+$.[7] I assume this cost function has a Cobb-Douglas form in which a reduction in type indicates a proportional increase in cost:

$$C(\cdot) = \theta^{-1} \exp\left(\sum_{m=1}^{M} \alpha_m s_m\right) q^{\beta_q} \ell^{\beta_\ell} \prod_{n=1}^{N} w_n^{\beta_n}. \tag{25}$$

[7] For regression variables and parameters, subscripts represent indexes.

Here, $\mathbf{s} \equiv (s_1, ..., s_M)'$ is a vector of publicly observable factors affecting cost. These variables include year and state fixed effects to control for region-wide annual shocks and state-level factors unlikely to have changed significantly in the four-year period. In addition, since corn is the primary output of the region, $\mathbf{s}$ includes county-level mean corn yields as a proxy for local factors that affect production such as average soil quality and climate characteristics.

Let $q^* \equiv \arg\max_q \{pq - C(\mathbf{w}, q, \ell, \theta)\}$ for a given output price $p \in \Re_{++}$. Algebraic manipulation of the first order condition for an interior solution to this profit maximization problem yields the following equation for the ratio of revenue to cost:

$$\frac{pq^*}{C(\mathbf{w}, q, \ell, \theta)} = \frac{\partial \ln C(\mathbf{w}, q, \ell, 1)}{\partial \ln q}. \tag{26}$$

Note that ratio (26) is independent of θ.

Following Diewert (1982), Shephard's Lemma and Eq. (26) provide a system of equations for estimating $C(\cdot)$. For this specification, the system of estimating equations for a typical observation is:

$$\ln \frac{C}{w_N} = \sum_{m=1}^{M} \alpha_m s_m + \sum_{n=1}^{N-1} \beta_n \ln \frac{w_n}{w_{N\,n}} + \beta_q \ln q^* + \beta_\ell \ln \ell + v_0 - \ln \theta \tag{27}$$

$$\frac{w_n {x_n}^*}{C} = \beta_n + v_n\ , \qquad n = 1, ..., N-1 \tag{28}$$

$$\frac{pq^*}{C} = \beta_q + v_N\ , \tag{29}$$

where C is observed cost. The vector of random noise for an individual landowner is $\mathbf{v} \equiv (v_0, v_1, ..., v_N)'$. Normalization of the cost function by w_N imposes positive linear homogeneity in input prices.

Since output is endogenous under the assumption of profit maximization, the state-level annual output price index p acts as an instrument for q^*. In addition, during the period in which the data were collected (1997-2000), the U.S. government was implementing a land set aside program (the Conservation Reserve Program). Since farmers were paid to remove land from production, the amount of land reported as being cultivated was likely correlated with individuals unobserved productivity characteristics. I therefore employ county population density as an instrument for ℓ. Finally, as noted by Fuss (1977), individual input price indices are likely to be endogenous since the component weights (e.g., relative use of diesel versus gasoline in determining an aggregate energy price) may also be correlated with unobserved landowner characteristics. To correct for this, I use state-level price indices as instruments for individual-level prices. The vector $\mathbf{z}$ denotes the exogenous variables for an individual landowner.

Since θ is unobservable to the government and the econometrician, $\ln\theta$ is modeled as a component of the error term in Eq. (27).[8] Assume that $\mathbf{v}$ and θ satisfy the following moment conditions:

M1. $E[\mathbf{v}|\mathbf{z}] = \mathbf{0}$;

M2. $E[v_0^3] = 0$;

M3. $E[\ln\theta|\mathbf{z}] = -\sigma\sqrt{2/\pi}$;

M4. $E\left[(\ln\theta - E[\ln\theta])^3\right] = \sigma^3(1-4/\pi)\sqrt{2/\pi}$.

Under M1, $\mathbf{v}$ is a mean-zero disturbance vector uncorrelated with the instruments. In addition, M2 states that the disturbance for Eq. (27) is symmetrically distributed. Assumptions M3 and M4 require that $\ln\theta$ be uncorrelated with the instruments, and indicate that it belongs to a "half-normal" distribution ($\ln\theta$ has the same distribution as $-|h|$, where h is a random variable distributed $N(0,\sigma^2)$). Thus, the distribution of $\ln\theta$ has a single parameter,

[8] Type cancels out of Eqs. (28) and (29).

σ, that determines all its moments.

As noted by Aigner et al. (1977), these distributional assumptions have two practical implications. First, consider a least squares estimator that ignores the type-dependent component of the error structure, using the incorrect moment conditions $E\left[v_0 - \ln\theta|\mathbf{z}\right] = 0$. By treating the expected compound error as mean zero, rather than mean $\sigma\sqrt{2/\pi}$, this regression upwardly biases estimates of coefficients corresponding to the intercept.[9] All other parameter estimates remain consistent. The second practical implication is that the compound error term $(v_0 - \ln\theta)$ is positively skewed with third central moment equal to $-\sigma^3\left(1-4/\pi\right)\sqrt{2/\pi}$.

The assumption of a half-normal distribution for the error component $\ln\theta$ is common in the stochastic frontier literature (Kumbhakar and Lovell, 2000). Intuitively, it can be justified by the notion that in a competitive economy the mode of the distribution of firms should be near the frontier. Since type is unobservable, however, it is impossible to formally test this structural hypothesis with cross-sectional data. Nonetheless, as depicted in Figure 1, the distribution of county-level average corn yields in the region under study appears to support the plausibility of this assumption.

Let $\mathbf{e} \equiv (e_1, ..., e_J)'$ denote the regression residuals for Eq. (27), where $j = 1, .., J$ indexes each observation. The third moment of the residuals is a consistent estimator for the third moment of the combined error term $v_0 - \ln\theta$. This suggests an additional equation that can

[9] Consider the special case in which $\mathbf{s}$ contains only a constant term. Least-squares methods can be thought of as estimating the following regression:

$$\ln C - \ln w_N = \sigma\sqrt{2/\pi} + \alpha + \sum_{n=1}^{N-1} \beta_n \ln\tilde{w}_n + \beta_q \ln q^* + \beta_\ell \ln\ell + \left[v_0 - \ln\theta - \sigma\sqrt{2/\pi}\right].$$

Note that the expected value of the error term in brackets is zero, but the presence of the (unobserved) first term on the right hand side will bias upwards the estimate of α.

be estimated sequentially after Eqs. (27)-(29):

$$e_j^3 = -\sigma^3 \left(1 - 4/\pi\right) \sqrt{2/\pi} + v_{j(N+1)}, \tag{30}$$

where $v_{j(N+1)}$ is random noise for observation j. The cube root of the estimate $\hat{\sigma}^3$ can then be used to correct the initial bias in the intercept.

Using these results, estimation of the system proceeds in three steps. The first step consists of ignoring $\ln \theta$, and estimating Eqs. (27)-(29) by system two-stage least squares (2SLS). Let $\mathbf{b}$ denote the vector of parameters for Eqs. (27)-(29). The system 2SLS estimator $\hat{\mathbf{b}}^{2SLS}$ is

$$\hat{\mathbf{b}}^{2SLS} = \left(\mathbf{Z}'\mathbf{X}\right)^{-1} \left(\mathbf{Z}'\mathbf{Y}\right), \tag{31}$$

where $\mathbf{X}$, $\mathbf{Y}$, and $\mathbf{Z}$ are respectively the equation-by-equation stacked right-hand side, left-hand side, and exogenous variables for all observations for Eqs. (27)-(29). This estimator is consistent for all parameters, except the intercept. The estimator $\hat{\mathbf{b}}^{2SLS}$ is likely to be inefficient, however, and generate inconsistent estimates of the covariance matrix. In addition to correlation of errors for the same observation across equations, the noise component may be heteroskedastic or influenced by unobserved shocks commonly affecting all landowners in the same geographic area. Such shocks may be short-lived or persist across time.

The next step addresses these potential problems. Following Pepper (2002), Wooldridge (2002), and Wooldridge (2003), the 2SLS residuals are used to construct a robust GMM estimator. The GMM estimator of the parameter vector $\mathbf{b}$ from Eqs. (27)-(29) can be expressed as:

$$\hat{\mathbf{b}}^{GMM} = \left(\mathbf{X}'\mathbf{Z}\mathbf{\Psi}^{-1}\mathbf{Z}'\mathbf{X}\right)^{-1} \left(\mathbf{X}'\mathbf{Z}\mathbf{\Psi}^{-1}\mathbf{Z}'\mathbf{Y}\right). \tag{32}$$

The weighting matrix $\mathbf{\Psi}$ is a function of the system 2SLS residuals:

$$\mathbf{\Psi} = \frac{1}{J}\sum_{i=1}^{I} \mathbf{Z}^{i\prime}\left(\mathbf{Y}^i - \mathbf{X}^i\hat{\mathbf{b}}^{2SLS}\right)\left(\mathbf{Y}^i - \mathbf{X}^i\hat{\mathbf{b}}^{2SLS}\right)'\mathbf{Z}^i, \tag{33}$$

where superscripts indicate that the respective matrices contain only the information for county i. Effectively, $\mathbf{\Psi}$ is the average of county-level weighting matrices calculated in turn from farm-level 2SLS residuals (for more details see Wooldridge (2002), pp. 328-330). For this estimator to be appropriate, I assume that errors from different counties are independent, and unobserved county effects are uncorrelated with the instruments. This estimator is asymptotically efficient in the presence of arbitrary heteroskedasticity and arbitrary county-level correlations both within and across time periods.

Finally, the third empirical moment of the GMM residuals from Eq. (27) is used as the left-hand side variable in Eq. (30). This last equation is then consistently estimated by ordinary least squares. The estimate $\hat{\sigma}^3$ is used to compute $\sigma\sqrt{2/\pi}$ and correct the bias in the GMM estimate for the intercept. Newey (1984) shows how the residuals from this sequence of regressions can be used to calculate an asymptotic covariance matrix for the entire system.

The consistency of estimates of the distribution of agent types and the intercept of the cost function (but not other parameters) depends upon both the half-normal distribution of $\ln\theta$ and the symmetry of $\mathbf{v}$ about its mean. Although these are arguably strong assumptions, they are in fact weaker than those typically employed in stochastic frontier analysis. The commonly-used maximum likelihood stochastic frontier approach (available for single equation estimation in several in econometric software packages) adds additional assumptions regarding the i.i.d. normality of $\mathbf{v}$. As noted by Kopp and Mullahy (1990), estimates of all parameters in such models are inconsistent if errors are not i.i.d. In addition, estimation of a complete stochastic frontier system using maximum-likelihood methods for non-i.i.d. errors is likely to be computationally unwieldy. To the author's knowledge this paper is the first to illustrate how the Pepper-Wooldridge-Newey GMM approach can be used to overcome these shortcomings in stochastic frontier analysis. The resulting estimator is not only robust, but computationally undemanding, even with a system of equations.[10]

3.2 Data and Estimation Results

The empirical analysis is conducted on a sample of midwestern U.S. agricultural producers. Producer cost and returns data come from 1997-2000 *Agricultural Resource Management Study* (ARMS) surveys conducted by the USDA's National Agricultural Statistics Service (NASS). The surveys are independent annual cross-sections in which it is not possible to track individual producers across time. ARMS contains data on input expenditures, output quantities, and land. Input and output price data come from the Bureau of Labor Statistics for capital and labor, the Federal Reserve for interest rates and NASS for other inputs as well as output commodities. County-level population densities come from the US Census Bureau website, and county-level corn yields come from the NASS website.

During the period in which the data were collected, the U.S. government implemented the Loan Deficiency Payment (LDP) program LDPs are loans made to farmers at harvest at a pre-determined rate per volume of output. In subsequent months, when farmers repaid the loan, they could do so either in full or at the prevailing market rate. LDPs thus acted as a price floor for program commodities. To account for this policy, I specify output price as the maximum of the loan rate or the market-year average price. The U.S. government also provided income support to farmers in the form of Production Flexibility Contracts and

[10] It is straightforward to program the estimation routine in a matrix language such as Gauss. Computations take seconds to complete on a standard PC.

Marketing Loss Assistance Payments. These payments were decoupled from output in the sense that they were made on the basis of area under cultivation, not production of a specific commodity. In the econometric model, I assume that these payments had no distortionary effect on production choice.

I aggregate outputs into a single category and variable inputs into capital services, labor, energy, and materials using a multilateral Tornqvist index (see Caves et al., 1982). Since ARMS surveys record capital assets as estimated market value at year end, I calculate capital services adapting the methodology of Hall and Jorgenson (1969).

The estimation procedure implicitly assumes all producers have the same general production technology (up to the type parameter). To limit possible specification bias, I focus attention on one relatively homogenous area, the "Heartland" Farm Resource Region.[11] This region comprises the corn belt. It includes the entire states of Illinois, Indiana, and Iowa, as well as portions of Kentucky, Minnesota, Missouri, Nebraska, Ohio, and South Dakota. It is the region with most farms, most cropland, and greatest value of production (Economic Research Service, 2000).

Table 1 presents summary statistics for the data. There are a total of 5,547 observations, where each observation corresponds to an individual farm in a single year. Without the inclusion of lump-sum government payments, the average return to non-land inputs is about 16 percent. As displayed in Figure 2, there is considerable variation in these net revenues, however. Although median returns are about $33 per acre, just over one third of the sample experienced negative market returns.

Table 2 reports the parameter estimates for this sample. For each endogenous variable (input price indices, land cultivated, and output), F tests reject at the 99 percent level the

[11] The USDA Economic Research Service divides the country into "Farm Resource Regions" with similar physiographic, soil, and climatic characteristics.

null hypothesis that excluded instruments do not have explanatory power.[12] I obtain the restricted profit function used in the theoretical model by simple algebraic manipulation of the estimated cost function and the first-order condition for q^*.[13] For the parameter estimates in Table 2, the restricted profit function satisfies all the theoretical restrictions imposed in Section 2. In addition, the estimated function is well-behaved, satisfying theoretical monotonicity and curvature conditions with respect to prices.

Although model specifications and data sets differ, input own-price elasticities are comparable to results from earlier studies of U.S. agriculture such as Ray (1982). Evaluated at the sample mean, the estimated average annual return to land is approximately \$57 per acre, without including any federal income support payments. For the three states entirely included in the sample (Illinois, Indiana, and Iowa), these payments averaged about \$45 per acre (Environmental Working Group, 2005). Including income support payments brings the average returns to about \$102 per acre. This figure is reasonably close to the average commercial rate of \$110 per acre paid by farmers who rented land for crop production in these three states (National Agricultural Statistics Service, 2001).

As shown in the previous section, empirical calibration of the optimal contract schedule requires two components: a restricted profit function for each type of producer and a probability distribution for type. The procedures described in this section provide precisely this information. The estimated cost function is used to calculate $r^i(a, \theta)$. The estimate $\hat{\sigma}^3$ is used to parameterize $f(\theta)$.

[12] The fact that the instruments are strongly correlated with the endogenous variables reduces the impact of inconsistency arising from a possible weak correlation of instruments with errors (see Bound et al., 1995).

[13] The functional form of the corresponding restricted profit function is:

$$r^i(\cdot) = \left[\beta_q - 1\right] \cdot \left[\theta^{-1} \left(\frac{\beta_q}{p}\right)^{\beta_q} \exp\left(\sum_{m=1}^{M} \alpha_m s_m\right) \prod_{n=1}^{N} w_n^{\beta_n} \left[\bar{a}^i - a\right]^{\beta_a}\right]^{\frac{1}{1-\beta_q}}$$

4 Policy Simulations

As an illustration of the usefulness of the methodological techniques in the previous section, I combine the parameter estimates with the theoretical results from Section 2. I use this information to calibrate simulations of the four hypothetical policy options to retire land in the midwestern United States. The simulations evaluate three policy decisions. The first two decisions involve the value of policy reform. I evaluate the benefits of changing from a single Pigouvian subsidy to a system in which the government acts as a monopsonist employing third-degree price discrimination, i.e., using a system of linear subsidies that vary by county. Next, I examine the benefits obtainable by shifting from a county-level Pigouvian subsidy to the second-best program. This comparison indicates the maximum amount the government should be willing to incur to develop an optimal policy without collecting additional information. The final comparison calculates the value of removing the information asymmetry. Suppose type were completely embodied in a measurable soil quality index. By comparing of the cost of the first and second best mechanisms, we obtain the maximum amount the government should be willing to pay to collect the soil quality information, i.e., to change from second to first-degree price discrimination.

For the simulation, I assume all farms in a given county are the same size (equal to the average), and all acreage is eligible for enrollment. Available land ranges from 74 to 1,150 acres per farm. County farm numbers are calculated by dividing total cropland in each county by average farm size. The data for these calculations come from National Agricultural Statistics Service (2002). Since about 5 percent of cropland in the Heartland region participates in the CRP, I set the enrollment target at that amount (about 7 million acres).[14]

[14]Information on regional cropland and CRP participation comes from Vesterby (2002) and Economic Research Service (2003).

For each policy, the numerical solution of the relevant necessary conditions characterizes the contract terms.[15]

Table 3 summarizes the simulation results. As discussed in Section 2, since they become progressively less constrained the programs must be non-increasing in cost as they go from the single Pigouvian subsidy to first-degree price discrimination. Interestingly, the only sizable difference among the programs is between second and first-degree price discrimination. The non-discriminatory subsidy required to attain the set-aside target is about 7.8 cents per acre, resulting in a total cost of $558 thousand.[16] Figure 3 illustrates the optimal distribution of Pigouvian subsidies when the government can condition on observable county characteristics. These subsidies range from 6.2 to 9.2 cents per acre, with a large plurality falling between 7.75 and 8.25 cents. Relative to a single Pigouvian subsidy, this third-degree price discrimination allows only modest cost savings of 0.3 percent, about $1,500 per year.

Using the Baron and Myerson mechanism, the lowest the government can spend to attain the target by using the information currently at its disposal is about $529 thousand, or 7.5 cents per acre. These savings represent about $24 thousand per year relative to the non-discriminatory subsidy, about 4.3 percent of the total cost.

The relatively small reductions in cost obtainable by engaging in second or third degree price discrimination contrast sharply with the first best. If the regulator could use information on individual types to engage in first-degree price discrimination, the cost of the program drops to about $193 thousand, or 2.7 cents per acre. The full information program is 63.6 percent less expensive than the second best. The maximum the regulator should

[15] All computations are performed in GAUSS.

[16] One reason the payments reported here are so low is that they do not include compensation for lost lump-sum income support payments. These payments are not private information to landowners. In principle, therefore, the government could compensate landowners for them exactly. Such compensation would increase the cost of the land set aside program. However, it would be offset by corresponding reductions in income support payments, resulting in no net increase in overall government expenditures.

be willing to pay to obtain the information necessary for first-degree price discrimination is \$336.4 thousand per year.

Rather than present the contract terms for each county, as an illustration Figure 4 depicts allocations for a second-best contract for a hypothetical county with mean characteristics. For comparison, it also illustrates contract allocations for first-degree price discrimination and a non-discriminatory Pigouvian subsidy, assuming all counties have identical characteristics.

The dotted line indicates contract terms for the Pigouvian subsidy. As indicated in panel (a), all contracts receive the same payment per acre. In panel (b), the enrollment rates by type for the Pigouvian subsidy are not visible since they coincide with those for the first-best contract. Enrollment decreases as type increases. This must be the case since each producer chooses to enroll a quantity of land such that his marginal opportunity cost is equal to the Pigouvian subsidy, and this marginal cost is increasing in type.

The solid line in both panels indicates contract terms for the first-best policy. The fundamental difference with the Pigouvian subsidy is that the government can fully eliminate any surplus payments. Farms with higher opportunity costs idle fewer acres. Thus, as shown in panel (a), smaller land set asides are matched with larger payments per acre. Such an allocation is not feasible if opportunity costs are private information since farms with lower costs could profitably mimic higher types by choosing contracts with low enrollment and high payments per acre.

The dashed lines depict the contract terms for the second-best policy. Panel (a) shows that the second-best payment schedule is non-linear, offering greater payments per acre as enrollment increases. The distortionary impact of the information asymmetry on the distribution of idled land across types can be seen in panel (b). Relative to the first best and

Pigouvian subsidy, the allocation is shifted so that lower types enroll more land, and higher types enroll less land. This distortion, combined with the fact that higher types receive lower payments per acre helps reduce the incentive for lower types to falsely claim to be high types. It is these distortions that reduce surplus payments relative to the Pigouvian subsidy. The optimal distortion is quite small, however, resulting in cost savings that are low in magnitude.

5 Conclusion

Recent articles have shown the usefulness of modeling type as an unobserved random variable for analysis of regulation under asymmetric information. Here, I extend earlier results in two directions. First, I extend the stochastic frontier literature by developing a GMM-based methodology for estimating a stochastic cost frontier for a profit-maximizing producer. This approach differs from earlier techniques in that it easily accommodates a system of equations (in this case a cost equation, expenditure share equations, and the ratio of revenue to cost) and is robust to arbitrary cross-equation correlation, heteroskedasticity, and geographic clustering. Further, it is computationally simple as it does not require non-linear optimization.

Second, I extend the empirical contract theory literature by using the empirical results to calibrate the theoretically optimal contract. Although the econometrician cannot directly observe producer type, the stochastic frontier approach permits consistent estimation of the technology and probability distribution of types in the population. This technique thus provides the necessary ingredients for specifying an optimal contract mechanism.

Application of this methodology to a simulate a voluntary environmental program to retire agricultural land yields some interesting results. The simulation permits comparison of the costs of four hypothetical programs that vary by degree of price discrimination. This type of analysis can provide guidance to policy makers interested in reducing the cost of voluntary environmental policies. For example, if the second-best program closely approximates the full-information program, reform efforts may be well spent in designing a complex system of non-linear contracts. For the sample of agricultural producers considered here, however, such does not appear to be the case. Simulated cost advantages from using second or third degree price discrimination are quite small. To the extent that there are differences in policy implementation costs (not explicitly modelled here), there may be reason to use a simple non-discriminatory policy like a uniform Pigouvian subsidy. Otherwise, the only manner of achieving significant cost savings is to try to overcome the information problem altogether by collecting detailed farm-specific agronomic data in an attempt to achieve first-degree price discrimination.

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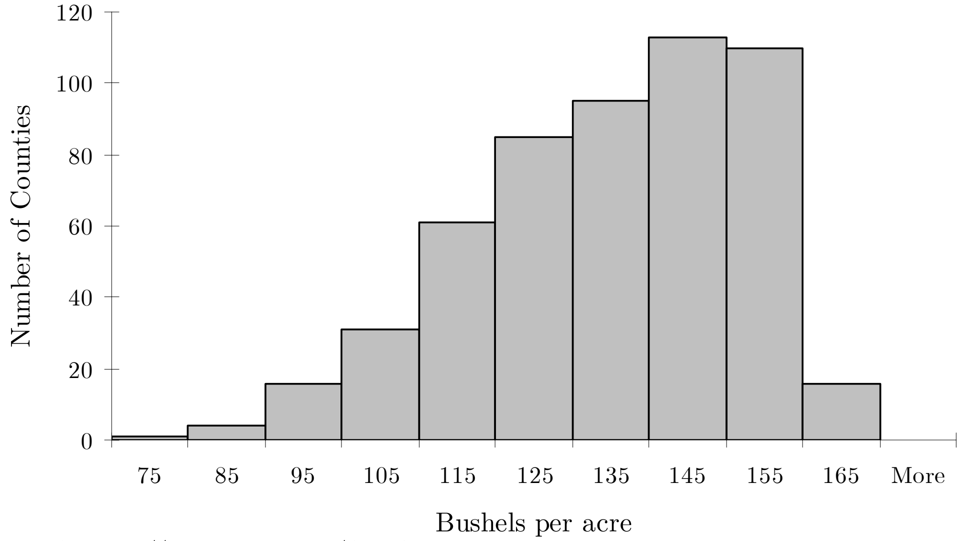
Figure 1: County Distribution of Corn Yields in “Heartland” Farm Resource Region



Data from http://www.nass.usda.gov/QuickStats

Figure 2: Sample Distribution of Net Returns per Acre

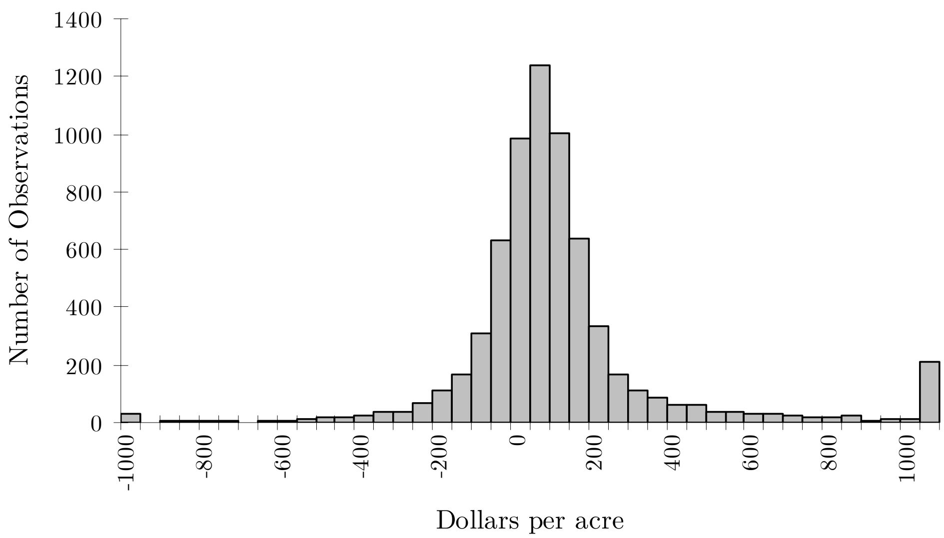


Table 1: Summary Statistics

	Mean	Standard deviation
Cost in current dollars (x1,000)	440.27	2,859.5
Input expenditure shares		
Capital	0.3760	0.1477
Labor	0.1776	0.1114
Energy	0.0423	0.0274
Materials	0.4040	0.1604
Revenue/cost	1.1615	0.8027
Output index (x1,000)	517.51	2,595.7
Acres	1039.2	1034.0
Price indices (farm level)		
Capital	0.9857	0.0335
Labor	0.9880	0.0843
Energy	0.9721	0.1216
Materials	0.9859	0.0357
Price indices (state level)		
Output	1.0102	0.1014
Capital	0.9856	0.0327
Labor	0.9880	0.0821
Energy	0.9721	0.1202
Materials	0.9857	0.0306
Cty. avg. yield	135.91	15.270
Cty. avg. population density	78.75	146.80
Number of observations	5,547	

Table 2: Parameter Estimates

Variable	Value	Standard error
ln Cty. avg. yield	-0.4686	0.0504
ln Capital Price	0.3622	0.0029
ln Labor Price	0.1811	0.0043
ln Energy Price	0.0429	0.0026
ln Materials Price	0.4138	0.0027
ln Output	1.1264	0.0029
ln Acres	-0.0807	0.0379
σ	0.8612	0.1112
R^2	0.8393	

Notes: Estimates robust to arbitrary heteroskedasticity and county clustering. R^2 for cost equation only.

Table 3: Simulation Results

Degree of Price Discrimination	Total Cost	Surplus Payments
First	192.9	0
Second	529.2	336.4
Third	551.8	358.9
None	553.2	360.4

Figure 3: Distribution of County-level Pigouvian Subsidies

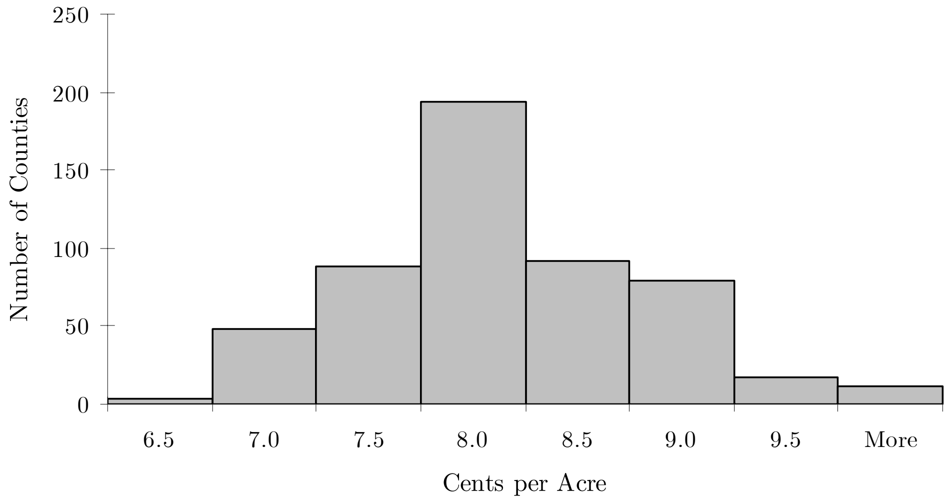
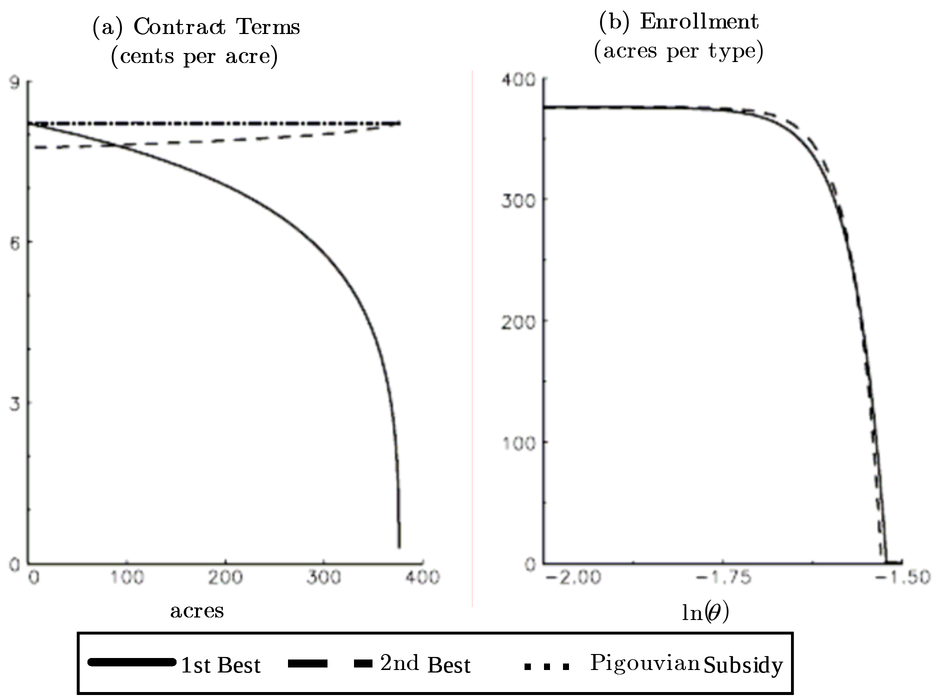


Figure 4: Regional First Best, Second Best, and Pigouvian Allocations



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